THERE'S STILL MADNESS IN THE METHOD: TESTING FOR FOOD MARKET INTEGRATION REVISITED

by

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Abstract

Food market integration, a situation in which arbitrage causes food prices in different markets to move together, is a precondition for the effectiveness of the market liberalisation programmes being implemented in many developing countries. Yet the standard econometric procedures which food price analysts use to test for the presence of market integration are both theoretically and statistically flawed.

This article evaluates the statistical performance of four commonly used econometric tests for market integration: the Law of One Price, the Ravallion Model, cointegration and Granger causality. A simulation model, which mimics many characteristics of developing country food markets, is used to generate artificial food price time series for both integrated and independent markets. Monte Carlo experiments using these food price series show that all four of the conventional tests for market integration are statistically flawed.

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I. INTRODUCTION

Liberalisation of food crop marketing is a common component of the structural adjustment and market reform programs being implemented in many developing countries. The leading international development agencies view such liberalisation as a means of promoting both static allocative efficiency and long-term agricultural growth (World Bank 1981; Elz 1987). However, in the absence of market integration (a situation in which prices in different markets move together if their price differential equals transfer costs) producers and consumers of agricultural products will not realise the gains from liberalisation. For in these circumstances, the correct price signals will not be transmitted down the marketing chain, and farmers will fail to specialise according to comparative advantage. In short, if 'getting prices right' (Timmer 1986) is seen as the crucial policy prescription for agricultural development, the presence of market integration is a vital precondition for it to be effective.

The growing trend towards the liberalisation of agricultural markets has therefore been accompanied by a veritable explosion in the literature on testing for food market integration[1]. The new test approaches that have been proposed - involving autoregressive distributive lags, cointegration, error-correction models and Granger causality - read like the table of contents of a recent textbook on time series econometrics. Yet the increasing complexity and sophistication of these methods has failed to challenge the basis of the traditional approach of testing for market integration using bi-variate correlation coefficients: that market integration can be discerned by assessing the underlying co-movement of prices. By failing to recognise the pivotal role played by transfer costs and the spatial arbitrage conditions, the literature has developed a series of tests that still provide a crude and inaccurate picture of the extent of integration in most developing country food markets. There remains as much madness in the methods used for testing market integration today, as when Harriss (1979) wrote her classic article a decade and a half ago.

This article assesses the robustness of four of the most commonly used econometric tests for market integration in a competitive multi-market trading situation. This is modelled using a spatial price equilibrium (SPE) model in the point-space tradition of Samuelson (1954), Takayama and Judge (1964), that is subject to both production shocks and general price inflation. Price data generated by this model is used to perform a series of Monte Carlo experiments in which the statistical reliability of each of the four econometric test procedures

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[1] See amongst others, Alexander and Wyeth (1994), Dercon (1995), Goletti *et al.,* (1995), Gordon (1994), Gupta and Mueller (1982), Heytens (1986), Palaskas and Harriss-White (1993), and Ravallion (1986)

can be evaluated objectively. The discontinuous patterns of trade generated by arbitrage in the presence of production shocks and transfer costs, are shown to undermine the concept of the perfect co-movement of prices on which the Law of One Price and the Ravallion model are predicated. In contrast, the cointegration and Granger causality approaches are found to be unable to distinguish between integrated and independent markets when both are subject to common and exogenous general price inflation.

II. MARKET INTEGRATION, TRANSFER COSTS AND THE CHARACTERISTICS OF SPATIALLY INTEGRATED FOOD MARKETS

Three forms of market integration are usually distinguished - integration across space (which is the focus of this paper), integration across time, and integration across product form (Timmer *et al.*, 1983; Tomek and Robinson 1990). In the case of spatial integration, two product markets are said to be integrated if, when trade takes place between them, price in the importing market equals price in the exporting market plus the transportation and other transfer costs of moving the product between the two markets. Put differently, if P_t^i denotes the price of food in the exporting market in period t, P_t^j denotes the contemporaneous price of food in the importing market and K_t^{ij} denotes transfer costs in the same period, then whenever:

$$P_t^i + K_t^{ij} = P_t^j \qquad \textbf{(1)}$$

trade occurs. But if:

$$P_t^i + K_t^{ij} < P_t^j \qquad \textbf{(2)}$$

then there is no incentive to trade. Similarly, markets are integrated across time when the expected price differential does not exceed the physical and financial costs of storage, and across product form when the price differential between two related commodities does not exceed processing costs.[2] Equations (1) and (2) are known in the literature as the spatial arbitrage conditions, and both are consistent with food market integration.

[2] In the case of spatial integration, the transformation process (transportation) is usually reversible and symmetric but in the other two cases, the transformation process will be asymmetric because reverse storage and processing are impossible.

Central to all forms of market integration are transfer costs, comprising transportation, storage and processing charges plus a modest allowance for trader's normal profit. Transfer costs determine the 'parity bounds' within which the prices of a commodity in two markets can vary independently of one another.[3] If markets are integrated, the price differential or spread between markets cannot exceed transfer costs. The arbitrage activities of traders, who ship a commodity between low and high price locations will raise price in some markets whilst lowering them in others until price differentials equals transfer costs and all opportunities for earning excess trading profits have been exhausted.[4]

The presence of transfer costs has three important implications for the characteristics of integrated food markets. First, since trade will only occur when price differentials equal or exceed transfer costs, if production or consumption is subject to shocks (due for example to the effect of weather on harvests or to shifts in demand), trade flows may be discontinuous between time periods. Under stable (non-stochastic) demand and supply conditions, regular trading patterns between markets will emerge based on stable price differentials. But if either production or consumption is variable over time (stochastic), price differentials will not always be sufficient to offset transfer costs and traders must decide between which, if any, markets it is profitable to arbitrage. Indeed, if transfer costs are sufficiently high, two markets may be 'integrated' even though no trade occurs between them. Seasonality in production may also lead to discontinuities, and even reversals, in trade flows between food markets (Timmer 1974).

Second, integrated food markets will be characterised by the simultaneous determination of prices. When trade flows are discontinuous, traders switch procurement between alternative sources of supply in response to changing market conditions. This can be seen most easily in the context of the radial market structure - in which a central food deficit market imports from a number of surrounding food surplus markets - hypothesised by Ravallion (1986). In this system, a shortfall of production in one supplying market will cause higher prices in the central market and induce higher flows of food into the central market from the other supplying markets. Higher export drives prices up in these supplying markets, until a new higher set of equilibrium prices is established, in which the spatial arbitrage conditions hold

[3] In the international finance literature, the parity bounds are sometimes referred to as the 'gold points'.

[4] The activities of storers and processors similarly ensure that food prices are kept temporally in line and that the price of processed products does not exceed those of their constituent raw materials plus processing costs.

between all markets. Despite the fact that individual traders assess the profitability of arbitrage between pairs of markets, all prices are determined simultaneously within this system.

A third characteristic of integrated food markets is the simultaneous determination of trade and storage. Because transfer costs can be substantial for both the transportation and storage transformations, profit-maximising traders are often faced with a trade-off between arbitrage across space and arbitrage across time. Recent work (Williams and Wright 1991) based on a multi-period, two-location model solved using dynamic programming methods, has demonstrated that the interaction of trade and storage decisions usually reduces price volatility and the frequency of trade. Trade flows are both lower in magnitude and more likely to be discontinuous for storable than non-storable commodities. Storage also induces positive autocorrelation in time series of prices and the volume of trade flows.

Another characteristic of integrated food markets, which is not due to the existence of transfer costs, concerns the presence of common trends across markets due to synchronous inflation or production seasonality. Such common trends, which are usually due to similar patterns of inflation or production seasonality, may result in time series of food prices that are non-stationary in levels but stationary in first differences. Many other real world commodity prices and macro-economic variables share this statistical characteristic (Baffes 1991; Palaskas and Varangis 1991), which is known - rather confusingly for the market integration literature - as integration of order one and abbreviated as *I(1)* . In contrast, time series that are stationary in levels are known as integrated of order zero, or *I(0)*. The non-stationarity of time series is associated with the econometric problems of spurious and inconsistent regression (Granger and Newbold 1974; Adam 1993) which invalidates standard econometric procedures unless the series are also cointegrated.[5]

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[5] Two or more time-series are said to be cointegrated if each is individually non-stationary but there is a linear combination of them that is stationary.

III. CONVENTIONAL APPROACHES TO TESTING MARKET INTEGRATION

The issue of how to test for food market integration occupies a voluminous literature going back to Jones's (1968) and Lele's (1967) work on staple food prices in, respectively, Nigeria and India using bivariate correlation coefficients. Advances in time series econometrics over the last two decades have led to the development of models addressing some of the perceived weaknesses in the correlation coefficients approach - in particular the statistical problems associated with common trends and the non-stationarity of food prices. But none of these models has questioned the central premise of the correlation coefficients approach: that market integration can be discerned by assessing the underlying co-movement between pairs of nominal market prices. This section explores the problems inherent in four commonly used econometric approaches to testing market integration: the Law of One Price, the Ravallion model, Granger causality and cointegration. It concludes that the increasing complexity and sophistication of the econometric methods used has failed to provide a theoretically adequate test for food market integration.

In comparing these tests, two econometric issues need to be borne in mind. First, different versions of the various test approaches will be appropriate for stationary and non-stationary data. Since, as noted above, most food price series tend to be non-stationary in levels but stationary in first differences, it is usually necessary to transform the original price series to convert them from *I(1)* to *I(0)* series. The estimating equations presented below assume that first differencing is the appropriate transformation to do this, although in some circumstances detrending or other transformations may also be necessary. Second, the estimation method used is dependent on the underlying market structure. Estimation via ordinary least squares will only be valid if prices in one market are exogenously determined. In the more commonly assumed radial or point-to-point market structures, food prices are simultaneously determined. In these circumstances, instrumental variable, limited information maximum likelihood or system methods should be used to correct for simultaneous equation bias.

The Law of One Price (Richardson 1978) is a test for the integration of markets within a single data period and in its usual form involves the regression of the current price change in one market on a constant and the price changes in another market. If, as before, P_t^j and P_t^j denote prices for a homogenous commodity in markets *i* and *j* in period t, the following equation is estimated:

$$\Delta P_t^i = \beta \Delta P_t^j \qquad \textbf{(3)}$$

The null hypothesis that *$\beta = 1$* is then tested using a standard t-test. If the null hypothesis can be rejected, then so is the Law of One Price.[6]

The Ravallion (1986) model allows price adjustment between markets to take time, but nests within it a test for short-run market integration that is equivalent to the Law of One Price. For non-stationary prices, the following error correction mechanism, which relates the contemporaneous change in price in market i to the contemporaneous and past spatial price differentials and the lagged price in market *j*, is estimated:

$$\Delta P_t^i = (\alpha_1 - 1)(P_{t-1}^i - P_{t-1}^j) + \sum\nolimits_{k=1}^{n} \alpha_k (P_{t-k}^i - P_{t-k}^j) + \beta_0 \Delta P_t^j + \sum_{k=1}^{n-1} (\beta_0 - 1 + \sum_{l=1}^{j} (\alpha_1 + \beta_1) \Delta P_{t-1}^j) + (\beta_0 - 1 + \sum_{k=1}^{n} (\alpha_k + \beta_k)) P_{t-1}^j + + \varepsilon_t \qquad \textbf{(4)}$$

When markets are integrated in the long term, the expression involving the coefficients on the last term will equal zero, and it takes several time periods for price changes to be fully transmitted between markets. Imposing this homogeneity restriction on the model, one may test for short-run integration using the null hypothesis that *$(\alpha - 1) = -1$* and *$\beta_0 = 1$*. Failure to reject the null implies immediate and perfect co-movement of prices.

The Granger causality approach (Granger 1969; Gupta and Mueller 1982; Alexander and Wyeth 1994) is similar to the Ravallion model in employing an error correction mechanism to assess the extent to which current and past prices changes in one market explain price changes in another. The following equation is estimated:

$$\Delta P_t^i = \alpha_0 + \sum_{k=1}^{n} \alpha_k \Delta P_{t-k}^i + \beta P_{t-n-1}^i + \sum_{l=1}^{m} \gamma_1 \Delta P_{t-1}^j + \delta P_{t-m-1}^j + \varepsilon_t \qquad \textbf{(5)}$$

where the number of lags (*n* and *m*) to include are determined by the use of a suitable information criterion. Rejection of the null hypothesis that *$\gamma_l = 0$* for *$l = 1..m$* and *$\delta = 0$* indicates that prices in market *j* Granger-cause prices in market *i*. If prices in market *i* also Granger-cause prices in market *j*, this indicates that prices are simultaneously determined.

Finally, the cointegration approach (Palaskas and Harriss-White 1993; Alexander and Wyeth 1994) measures whether two markets are integrated in the long term by assessing whether

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[6] The Law of One Price can be estimated using either differences of the original series or differences in their natural logarithms. When differences of the original price series are used, absolute marketing margins become a maintained hypothesis. When log differences are used, proportional marketing margins are implicitly assumed (Dercon 1995).

their prices wander within a fixed band. The usual two-step residual-based test, due to Engle and Granger (1987), involves the estimation of the following 'cointegrating regression' using ordinary least squares:[7]

$$P_t^i = a + bP_t^j + \varepsilon_t \quad \textbf{(6)}$$

The residuals from this cointegrating regression are then tested for stationarity using the augmented Dickey Fuller test (Dickey and Fuller 1979). This involves estimating:

$$\Delta\varepsilon_t = \alpha + \beta\varepsilon_{t-1} + \gamma_t + \sum_{j=2}^{N} \delta_j \Delta\varepsilon_{t-j} \quad \textbf{(7)}$$

where *t* is a time trend. If the null hypothesis that $\beta = 0$ can be rejected this indicates that the series are cointegrated.[8] Cointegration indicates that a long-run reduced form relationship exists between the two time series, but is neither a necessary nor a sufficient condition for market integration (Barrett 1996).[9]

The null hypotheses in these four test approaches divide into two groups. The Law of One Price and Ravallion models test whether price changes in market *i* will be translated on a one-for-one basis to market *j*, either instantaneously (the Law of One Price) or with lags (the Ravallion model). But prices in different markets will only move on a one-for-one basis if the intermarket price differential is equal to transfer costs. So for the Law of One Price and Ravallion model to be valid, the spatial arbitrage conditions must always bind, trade flows must occur in every period, and transfer costs must be constant. This is an extremely restrictive definition of market integration. In contrast, the Granger causality and cointegration approaches test for much more general notions of equilibrium between markets. These tests allow for price co-movement to be less than perfect, for trade flows to be discontinuous and for there to be variations in transfer costs. But, like the Law of One Price

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7 The cointegrating equation may always be estimated using ordinary least squares because of the 'super-consistency' property of cointegrated time series (Stock 1987).

8 As the Gauss-Markov theorem's assumption of finite variance of the error term is violated when series are *I(1)*, special critical values must be used in performing this test. These have been tabulated by Dickey and Fuller (1979), Engle and Yoo (1987) and MacKinnon (1990) and are always negative.

9 Cointegration is an unnecessary condition for market integration because if transfer costs are non-stationary, arbitrage between two markets may be efficient even though their price series are not cointegrated. Cointegration is an insufficient condition for market integration because two price series may be cointegrated while their price differential is too small to offset transfer costs. This point may be easily seen by considering two identical, non-stationary price series from two spatially separated markets between which arbitrage is costly. Since the price series are identical but non-stationary, they are cointegrated (in equation (6) a = 0 and b = 1), yet arbitrage between the two markets would never be profitable.

and Ravallion models, they also assume a linear relationship between market prices which is inconsistent with the discontinuities in trade implied by the spatial arbitrage conditions. There is also a tendency to assume that first differencing will remove all common trends in food price time series. The consequences of these two types of market integration tests within a multi-market trading system will now be explored.

IV. THE SPATIAL PRICE EQUILIBRIUM MODEL

In order to examine the statistical performance of the four conventional approaches to testing market integration described above, a model of spatial price equilibrium in the point-space tradition of Samuelson (1952) and Takayama and Judge (1964) was constructed.[10] This model generates artificial price series which mimic three of the four characteristics of integrated food markets described in the Section I. It also generates time series for a pair of independent markets which never trade. Repeated application of the conventional tests for market integration to the artificial price series generated by the model, allows the statistical reliability of the four tests to be assessed in a situation in which the underlying form of market integration is known by construction.

The spatial price equilibrium (SPE) model developed involves a single homogenous commodity (henceforth described as 'grain') traded between six discontinuous regions (which it is convenient to think of as 'islands'). The six islands all have populations of the same size and characteristics (so demand in all six islands is identical) but different amounts of arable land (so their levels of grain production differ). The demand for grain is assumed to be stable over time (non-stochastic) and is described by identical inverse demand functions in all six islands.[11] Mean grain is, however, stochastic so production levels differ between islands due to a series of uncorrelated harvest shocks. These occur once a period ('crop year') and

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[10] Samuelson's (1952) seminal article considered a *n* region economy for a non-storable commodity with a perfectly elastic supply of transportation services. He demonstrated that under perfect competition the prices and trade flows that meet traders' arbitrage conditions are identical to those resulting from maximisation of social surplus (where social surplus is defined as the sum of producer and consumer surpluses minus transportation costs). Takayama and Judge (1964) proved that, if demand and supply curves are linear, Samuelson's model could be solved using quadratic programming. Applications of the Samuelson-Takyama-Judge SPE model include those of Hall, Heady and Plessner (1968), Kennedy (1974), Nagy, Furtan and Kulshreshta (1980), Shapiro & White (1982) and Bivings (1992).

[11] Like production, the demand for grain will also often be stochastic due to the influence of 'shifters', such as weather or tastes. For the purposes of generating discontinuous trading patterns, it is sufficient to assume that either demand or supply is stochastic - and the latter assumption has been adopted here.

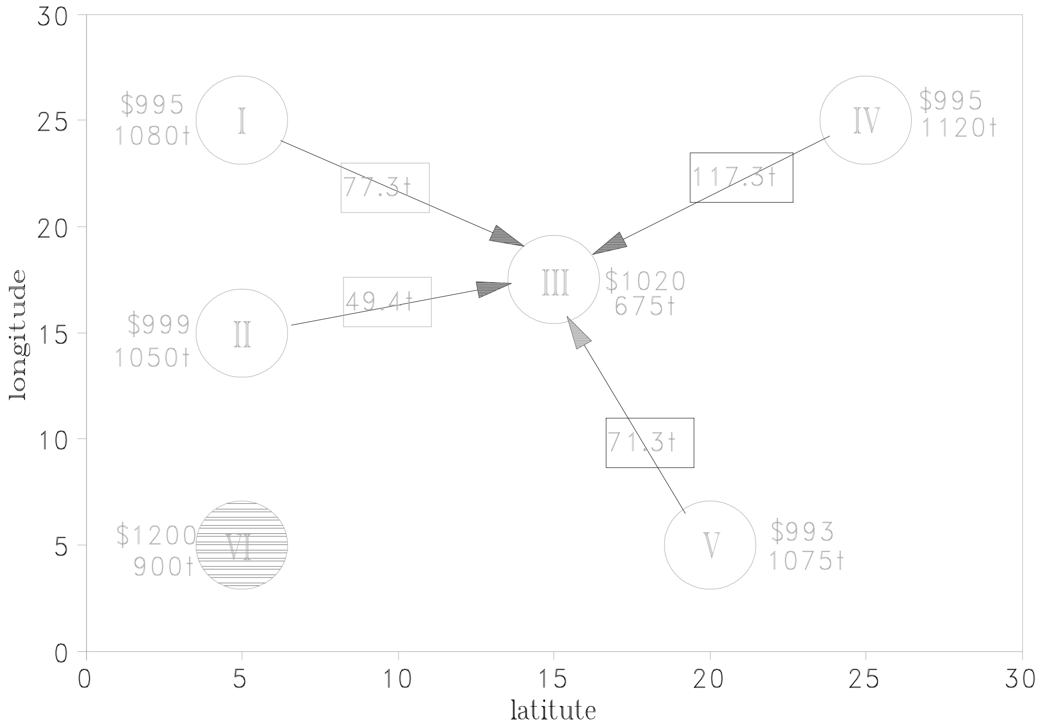
represent the influence of weather, pests and vectors on grain production in each of the islands. For simplicity, the harvest shocks are assumed to possess independent and identical normal distributions.

It is assumed that inter-island grain markets are competitive and that grain may be shipped between islands I to V in unlimited quantities. Island VI, however, pursues a policy of food self-sufficiency and never trades with any of the other five islands, although it would usually be profitable for her to do so. Shipping costs (which represent the only element in transfer costs) vary in direct proportion to the distance between islands and comprise between 2 and 5 per cent of the delivered grain price, depending on the distance between islands. Due to changing capacity utilisation in the shipping industry and other external factors (e.g., the weather, dockers' strikes), unit shipping costs do, however, vary a little from period to period.

Spatial price equilibrium is established in the model when all arbitrage opportunities between islands I to V have been exhausted. When there is trade between pairs of islands, the grain price in the importing island equals the grain price in the exporting island plus shipping costs. But when there is no trade between islands I to V, shipping costs exceed the spatial grain price differential. In the case of the autarkic Island VI, price differentials usually exceed transfer costs but it nevertheless does not engage in trade because of its policy of food self-sufficiency. Figure 1 shows the physical configuration assumed for the six-islands, together with their mean production levels and the resulting pattern of trade when unit shipping costs equal $2 per ton mile.[12] The similarity between this trading pattern and the radial trading pattern assumed by much of the market integration literature should be apparent.

[12] Note that the configuration of islands in Figure 1 precludes the need for transhipment. Given the location of the islands, it is always cheaper for two islands to trade with each other directly rather than by shipping through a third island. It is also never economic for an island to import and export grain simultaneously in this model.

Figure 1: Solution to the Spatial Price Equilibrium Model with Mean Production



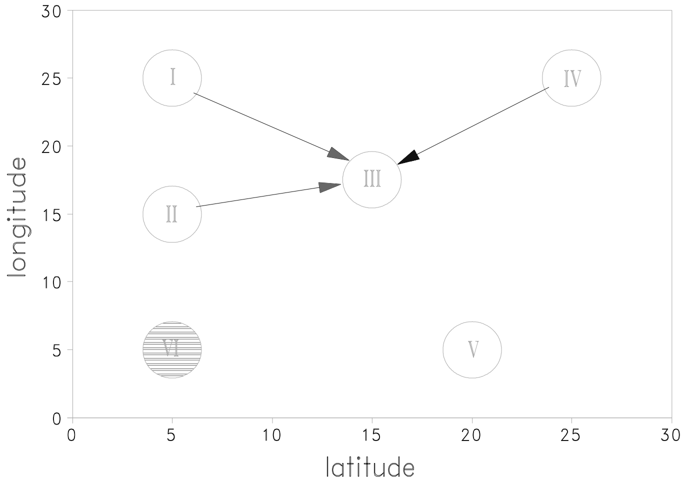
As grain production (and to a lesser extent shipping costs) varies from period to period, prices and the quantities of grain shipped between islands I to V adjust in response to changing arbitrage opportunities. So when the SPE model is simulated repeatedly, a sequence of spatial price equilibria is established, in which the spatial arbitrage conditions hold between islands I to V and grain markets clear in all six islands in every period. Associated with each equilibrium is a set of market clearing prices, trade flows and unit shipping costs.

With a standard deviation of harvest shocks of 50 tons and mean unit shipping costs of $2 per ton mile, the radial trading pattern shown in Figure 1 occurs in approximately 72 per cent of simulations of the SPE. In the remaining 28 per cent of simulations, trade deficit islands switch their grain procurement between alternative sources of supply in response to changing market conditions and other patterns of trade occur.[13] The two most common of these are shown in Figures 2a and b, which occur in 13 per cent and 7 per cent of simulations respectively. Four or five other trade patterns occur once or twice in every 1,000 simulations of the model.

Figure 2: Other Trade Patterns Generated by the Spatial Price Equilibrium Model

(a)

[13] Note that in this context 'trade pattern' refers to the direction, but not the volume of trade flows between islands.



I
II
III
IV
V
VI
longitude
latitude
0
5
10
15
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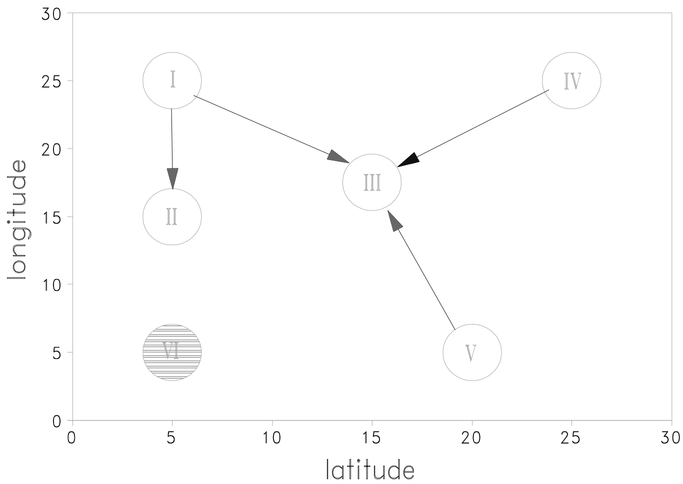
25
30

(b)



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To incorporate common trends and the non-stationarity of food prices into the SPE model, the prices it generated were adjusted by a common and exogenously determined inflationary process. This was modelled using the price index:

$$I_t = (1 + \mu + \sigma * \varepsilon_t) I_{t-1} \qquad \textbf{(8)}$$

where μ and σ are parameters representing the mean and standard deviation of inflation and ε_t is a random variable drawn from the standard normal distribution. Readers may find it helpful to think of the stationary price series generated by the SPE model as 'relative' prices which, when inflated by the price index, become non-stationary, *I(1)* nominal prices. For $\mu = 0.012$ and $\sigma = 0.012$, a typical realisation of this process, together with two of the relative and nominal price series from the SPE model, are shown in Figure 3, overleaf.

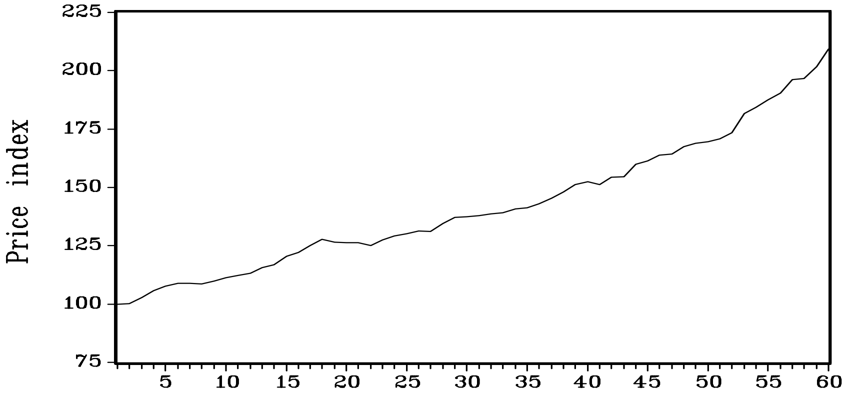
The trade flows, trading patterns and nominal food series that are generated by the SPE model and inflationary price index possess three of four key characteristics of integrated food markets identified in the introduction (i.e. discontinuous trade, price endogeneity and the presence of common trends). The final characteristic of integrated food markets - the simultaneous determination of trade and storage - does not apply to the SPE model which, in effect, assumes grain to be non-storable.[14] However, many of the time series properties that are known to be induced by storage, such as the discontinuity of trade flows and positive autocorrelation in prices (Williams and Wright 1991), are already present within the model.

The algebraic formulation of the SPE model, together with its programming solution, is discussed in the Appendix.

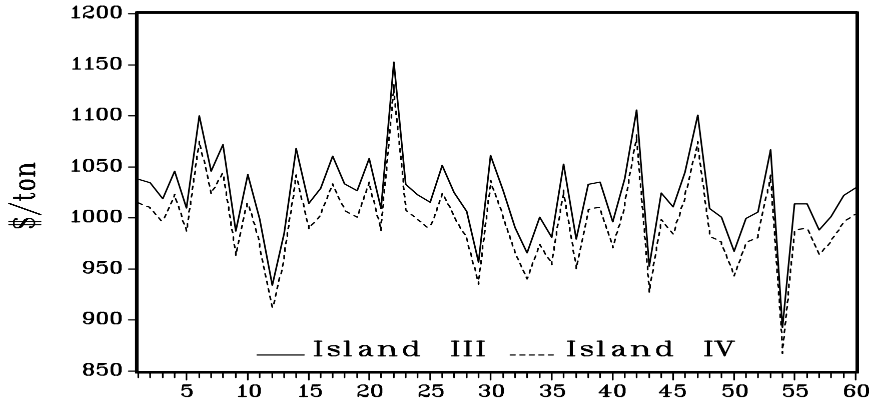
[14] Construction of a complete model of trade and storage for the six island SPE model considered here would be a formidable task, involving the solution of a five state variable dynamic programming problem. Solution of such a model is, probably, beyond the scope of current computing technology.

Figure 3: A Typical Realisation of the Inflationary Process and Two Price Series from the Spatial Price Equilibrium Model

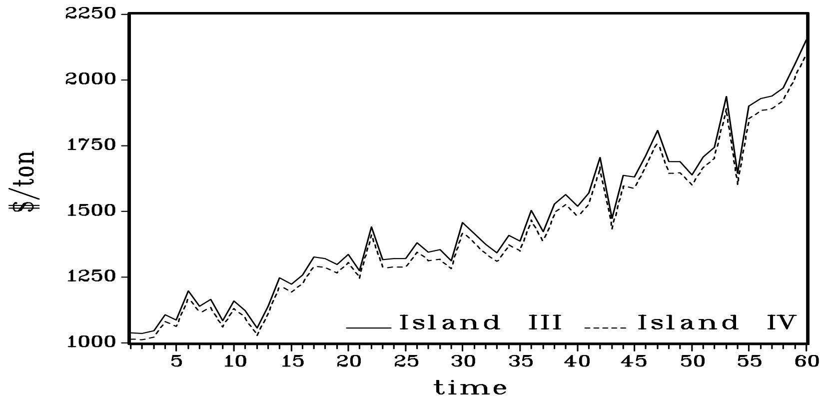
PRICE INDEX



RELATIVE PRICES



NOMINAL PRICES



V. MONTE CARLO RESULTS

To investigate the robustness of the four conventional tests of market integration, the SPE model was used to generate multiple time series of production, trade flows and relative prices typical of the short sample sizes available in most developing countries. A series of Monte Carlo experiments in which the conventional tests for market integration were replicated 1,000 times using sample sizes of 60 and 120 observations (i.e., five and ten years of monthly price data) were then conducted. These experiments allow the statistical reliability of each of the tests for market integration described in Section III to be evaluated, in a situation in which the underlying form of market integration and data generation mechanism are known with certainty.

Tables 1 and 2 show the Monte Carlo results from repeated application of the four conventional tests for market integration to price data from the SPE model.[15] Each entry in the tables corresponds to the percentage of replications in which the null hypothesis was rejected at the 5 per cent level of statistical significance. Two pairs of islands have been selected from the SPE model to represent the opposite scenarios of integrated and independent markets. Island II and III, which trade in 84 per cent of observations, represent the integrated markets case. Islands III and VI, which never trade, represent the independent markets case. Selection of alternative pairs of islands changes the tabulations of the hypothesis tests a little, but does not alter the overall conclusion of the Monte Carlo analysis. Results using both ordinary least squares (OLS) and instrumental variable (IV) estimation are presented in order to show the bias that results from failure to correct for the simultaneous determination of prices.

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[15] All econometric estimation was performed using Time Series Processor (TSP) version 4.2 (Hall, Cummins and Schnake 1991) in a UNIX SunOS 4.1.3 environment via a Sun SPARC2 workstation.

Table 1: Monte Carlo Results from the Spatial Price Equilibrium Model
(% rejections of the null hypothesis at the 5% level)

Null Hypothesis	Method of Estimation	Integrated Markets	Independent Markets
'Law of One Price'	OLS	20.3	99.9
	IVs	14.9	
Ravallion Model			
1. Long-run	OLS	100.0	9.9
Integration	IVs	9.4	
2. Short-run strong	OLS	100.0	99.9
form integration	IVs	90.5	
Not Cointegrated	OLS	99.8	99.5
Not Granger Causing	OLS	100.0	100.0
	IVs	94.8	

NOTES: 1) Based on 1,000 replications each of 120 observations.
2) In the shaded quadrants a **high** percentage of rejections would be expected if the tests were reliable. In the unshaded quadrants a **low** percentage of rejections (about 5 per cent) would be expected if the tests were reliable.
3) The null hypothesis of non-cointegration is only tested using ordinary least squares (OLS) because of the 'super-consistency' of cointegrated time series.
4) Instrumental variables (IVs) estimation is not used in the independent markets case because islands III and VI never trade.

In interpreting the results of the Monte Carlo analysis for the integrated markets case, it should be remembered that, by construction: (i) the spatial arbitrage conditions always hold between islands II and III; (ii) there are no lags in price adjustment (i.e., their markets are integrated instantaneously); and (iii) all nominal price series are *I(1)* and cointegrated, since their linear combination, $P_t^i - \beta P_t^j$, is *I(0)* for all island pairs. So assuming a 5 per cent level of statistical significance, one would expect the null hypothesis of the Law of One Price to be rejected in around 5 per cent of replications if the Law of One Price is a good test for market integration. Given the instantaneous nature of price transmission, Ravallion's tests for both long and short run integration should also be rejected in about 5 per cent of replications. In contrast, the null hypotheses of non-cointegration and non-Granger causality (which are inconsistent with market integration) should be rejected in 95 per cent of replications.[16] It is not, however, possible to specify how often the null hypothesis should be rejected in the independent markets case because there is an unspecified amount of Type II errors implicit in

[16] Testing for non-Granger causality is equivalent to Ravallion's test for market segmentation as long as contemporaneous price in the central market is included as a regressor.

their formulation.[17] Conventional statistical power calculations also cannot be performed, because the 'true' value of the underlying parameters is not known in the SPE model when production (or demand) is stochastic. The shaded quadrants in the tables, however, indicate the instances in which a high frequency of rejection of the null hypothesis would be expected.

It will be apparent from the test results in Table 1 that, with a sample size of 120 observations, only the tests for cointegration and Granger causality tests have acceptable size in the integrated markets case.[18] The Law of One Price's null hypothesis is rejected more than four times as often as it should be when estimated using OLS, and about three times as often as it should be with IV estimation. The null hypothesis of Ravallion model's tests for long and short-run integration are rejected uniformly when estimated by OLS. Under IV estimation (the method recommended by Ravallion), the test for long-run integration has improved size.[19] But in the case of short-run integration (strong-form), even IV results in rejection of the null hypothesis in more than 90 per cent of the time. In contrast, the cointegration and Granger causality approaches, reject the null hypothesis correctly in the vast majority of replications.

In the independent markets case, Table 1 shows that the null hypotheses of the Law of One Price and Ravallion's test for short-run integration are correctly rejected using OLS in 99.9 per cent of replications. The cointegration and Granger causality approaches however, show themselves to be unable to distinguish between integrated and independent markets. Their null hypotheses are rejected almost as often in the independent as the integrated markets case. (Note that it is inappropriate to apply IV estimation in the independent markets case because islands III and VI never trade.)

For the sake of comparison, Table 2 reports the results of repeating these Monte Carlo experiments with a sample size of 60 observations. One observes less rejections of the null hypothesis for the Law of One Price and Ravallion models when the sample size is decreased from 120 to 60 observations. The frequency with which the null hypotheses of non-

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[17] See Gordon (1994) for an expansion of this point for the Ravallion model.

[18] The size of a test is the probability of a Type I error (rejection of a true null). The power of a test is the probability of a Type II error (acceptance of a false null).

[19] Note that the performance of Ravallion's test for strong form integration is highly dependent on the restriction of long-run market integration being imposed on his model (by dropping lagged prices in island III in the error correction form of his model). When this restriction is not imposed, Ravallion's test for weak-form short-run integration is rejected in 100% of cases.

cointegration and non-Granger causality are rejected is not, however, greatly affected by this halving of the sample size. The contrast between the Law of One Price's and Ravallion models' poor performance in the integrated markets case, and the inability of the tests for cointegration and Granger causality to distinguish between integrated and independent markets, is unaffected.

Table 2: Monte Carlo Results from the Spatial Price Equilibrium Model (% rejections of the null hypothesis at the 5% level)

Null Hypothesis	Method of Estimation	Integrated Markets	Independent Markets
'Law of One Price'	OLS	13.4	99.4
	IVs	8.1	
Ravallion Model			
1. Long-run	OLS	100.0	6.9
Integration	IVs	4.0	
2. Short-run strong	OLS	100.0	99.4
form Integration	IVs	78.1	
Not Cointegrated	OLS	97.9	99.0
Not Granger	OLS	100.0	93.4
Causing	IVs	95.2	

NOTE: Based on 1,000 replications each of 60 observations (see also Notes 2, 3 and 4 in Table 1 above).

The principal explanation for the perverse and contrasting statistical performance of the tests in these Monte Carlo simulations lies in the nature of their null hypotheses. As noted in Section III, the Law of One Price and Ravallion model are essentially tests of whether there is perfect comovement of prices. But prices in the different islands of the SPE model will only move on a one-for-one basis if price differentials are equal to transfer costs. When they are not, prices will move independently. Since islands II and III only trade with each other in 84 per cent of observations, it is hardly surprising that the Law of One Price and Ravallion models frequently reject the one-for-one comovement of prices between them. The presence of production stocks and transfer costs in the SPE model creates discontinuous trade patterns, which undermines the perfect price comovement upon which the tests are based.

The cointegration and Granger causality approaches are unable to distinguish between the integrated and independent market cases for a different reason. Their null hypotheses test for the much more general notion of a long-run equilibrium between markets. Tests for cointegration find a long-run relationship for both integrated and independent markets not because of inter-island arbitrage, but due to the presence of the inflationary trend specified by equation (8). The price changes due to this common trend overwhelm relative price changes due to changes in the balance of demand and supply and to inter-island arbitrage (in fact, the underlying relative price differentials between islands III and VI, which never trade, are entirely random). Similarly, the Granger causality tests appear to show that prices in Island III 'Granger-cause' those in Island VI, when it is really the common inflationary trend that is driving price changes in both markets. In each case, first-differencing the prices series is insufficient to remove the common trend in the price series created by equation (8), despite the fact that they are shown to be stationary when the univariate augmented Dickey Fuller test is applied. This demonstrates the relatively low power that the augmented Dickey Fuller test has in distinguishing trend-stationary from difference-stationary processes in finite samples (Blough 1992). It also confirms Palaskas and Harriss-White's (1993) statement that cointegration is a necessary but not a sufficient condition for market integration.

In short, these Monte Carlo simulations demonstrate that conventional econometric tests for market integration catch the food price analyst on the horns of a dilemma: some tests, such as the Law of One Price and Ravallion model, are too strong to detect market integration while others, such as cointegration and Granger causality, are too weak.

VI. CONCLUSIONS

This article has considered the performance of four pairwise tests for market integration using a model of spatial price equilibrium subject to production shocks and general price inflation. Tests for market integration based on hypotheses tests of whether one or more regression coefficients sum to unity, such as the Law of One Price and Ravallion model, have been shown to perform poorly. This is because food prices will not move on a one-for-one basis if trade does not occur between market pairs in every period.

Due to the less restricted nature of their null hypotheses, tests for cointegration and Granger causality are more robust to such discontinuities in trade flows. By the same token, they are also less informative, as they are unable to distinguish between integrated and independent markets in the presence of a common trend. This points to the inappropriateness of using tests for cointegration and Granger causality as tests for market integration, despite their popularity in the agricultural, financial and futures markets literature (Alexander and Wyeth 1994; Palaskas and Harriss-White 1993; Taylor and Tonks 1989; Chowdhury 1991). Testing the stationarity and cointegration of price series should, nevertheless, remain crucial steps in the exploratory data analysis that precedes the modelling of food or agricultural market prices.

In short, just as in the fable, 'the Emperor has no clothes'. Except for one special case - the radical market model with continuous trade flow to the central market and constant transfer cost - it is impossible to test for food market integration adequately using price data alone.[20] A more reliable test for food market integration requires explicit consideration of both transfer costs and the simultaneous nature of price formation. One such test - which extends earlier work on stochastic frontier and switching regression models (Aigner, Lovell and Schmidt 1977; Spiller and Wood 1988; Sexton, Kling and Carman 1991) by using information on mean transfer costs in addition to nominal food prices - is the parity bounds model (Baulch 1995).

20 Similar conclusions have been reached from the analytical derivation of a conditional expection function for the Law of One Price by McNew (1994), and from an empirical examination of purchasing power parity in foreign exchange markets by Davutyan and Pippenger (1990).

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Appendix: Algebraic Formulation of the Programming Solution to the Spatial Price Equilibrium Model Under Stochastic Production

Inverse Demand

$P^i_t = \alpha + \beta D^i_t$ i = 1 to 6

Internal Supply

$H^i_t = \gamma^i + \varepsilon^i_t$ where $\varepsilon^i_t \sim N(0,\sigma^i)$ i = 1 to 6

Total Supply

$S^i_t = H^i_t + F^{ij}_t$ i,j = 1 to 5

$S^i_t = H^i_t$ i=6

Transfer Costs

$TC^{ij}_t = (1 + \nu_t)\kappa F^{ij}_t$ where $\nu_t \sim N(0,\sigma)$ i,j = 1 to 5

Equilibrium Conditions

$P^j_t \leq P^i_t + TC^{ij}_t$ if $F^{ij}_t \geq 0$ i,j = 1 to 5

$D^i_t = S^i_t = Q^i_t$ i = 1 to 6

where P^i_t denotes the price of grain in island i in period t, D^i_t is the quantity of grain demanded in i in period t, F^{ij}_t is the grain flow from island i to island j in period t and α, β, γ, κ and σ are parameters. Demand parameters (α = 3,000 and β = -2) were chosen that would lead to an average price of $1,000 per ton and average consumption of 1,000 tons of grain in each island in the absence of transportation costs.

Samuelson (1952) demonstrated that the social planner's problem of maximising social surplus:

$$Max\sum_{i=1}^{5}[(P_t^iQ_t^i)+(Q_t^i(\alpha-P_t^i))-\sum_{j=1}^{5}(\kappa F_t^{ij})]$$

$$s.t. \quad F_t^{ij} \geq 0,\ P_t^i \geq 0 \qquad \forall i,j$$

is equivalent to the competitive equilibrium solution to the problem that would be achieved by the arbitrage activities of a large number of producers and consumers. The spatial price equilibrium model can therefore be solved by maximisation of the objective function in equation (1) above. Furthermore, if demand and supply are linear, then the objective function to be maximised will be quadratic (Takayama and Judge 1964).

This standard problem in concave programming was solved using a FORTRAN 77 program employing the International Math and Statistic Library subroutine NCONG itself based on Schittkowski's (1986) NPPQL algorithm. ε_t^i and ν_t were simulated using the International Mathematics and Statistics Library subroutine DRNNOR, which generates pseudo-random numbers from a standard normal distribution using the inverse CDF method. Copies of this program are available from the author on request.